DOCUMENT RESUME

ED 377 205 TM 022 347

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TITLE Validating a Swedish Version of the Beck Depression Inventory for Nonclinical Adolescents: A Confirmatory Factor Analytic Study.
PUB DATE 94
NOTE 37p.; Paper presented at the International Congress of Applied Psychology (23rd, Madrid, Spain, 1994).
PUB TYPE Reports - Research/Technical (143) -- Speeches/Conference Papers (150)

EDRS PRICE MF01/PC02 Plus Postage.
DESCRIPTORS *Adolescents; Correlation; *Depression (Psychology); Foreign Countries; Goodness of Fit; High Schools; High School Students; Hypothesis Testing; *Measures (Individuals); Psychometrics; *Validity
IDENTIFIERS *Beck Depression Inventory; Confirmatory Factor Analysis; *Sweden

ABSTRACT

The intent of the present study was to test for the validity and equivalency of a second-order factorial structure of the Beck Depression Inventory for and across three independent samples (n1=661; n2=239; n3=196) of nonclinical Swedish adolescents; the model under study derived from a cross-validated study of Canadian high school adolescents (Byrne and Baron, 1993a). Model fitting, testing, and equating were based on the analysis of covariance structures within the framework of a confirmatory factor analytic model. Of major importance was the tenability of the hypothesized model for Swedish adolescents; only minor differences that involved correlated errors and one cross-loading (Sample 2) were evidenced. Results are expected to be of substantial interest to both researchers and clinicians whose concerns focus on depression as it bears on this population. (Contains 39 references, 1 table, and 4 figures.) (Author)

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ED 377 205

Validating a Swedish Version of the Beck Depression Inventory for Nonclinical Adolescents: A Confirmatory Factor Analytic Study

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Paper presented at the 23rd International Congress of Applied Psychology, Madrid 1994

TM022347

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Abstract

The intent of the present study was to test for the validity and equivalency of a 2nd-order factorial structure of the Beck Depression Inventory for and across three independent samples (n_1=661; n_2=239; n_3=196) of nonclinical Swedish adolescents; the model under study derived from a cross-validated study of Canadian high school adolescents (Byrne & Baron, 1993a). Model fitting, testing, and equating were based on the analysis of covariance structures within the framework of a confirmatory factor analytic model. Of major importance was the tenability of the hypothesized model for Swedish adolescents; only minor differences that involved correlated errors and one cross-loading (Sample 2) were evidenced. Results are expected to be of substantial interest to both researchers and clinicians whose concerns focus on depression as it bears on this population.

The Beck Depression Inventory: Testing and Cross-validating a Second-order Factorial Structure for Swedish Nonclinical Adolescents

Despite its original development for use with clinical samples, the Beck Depression Inventory (BDI; Beck, Ward, Mendelson, Mock, & Erbaugh, 1961) is now the most widely used measure of depression for nonclinical adolescents. Addressing a need to validate the factorial structure of the instrument for use with this population, Byrne and Baron (1993a, 1993b; Byrne, Baron, & Campbell, 1993, 1994) conducted a series of studies to investigate the psychometric soundness of such application for English- and French-speaking Canadian adolescents. The present study extends this work in testing for the validity and replication of an hierarchical factorial structure of the BDI for Swedish nonclinical adolescents.

Given the popularity of the BDI for use with high school adolescents, it is surprising to find only a modicum of studies that have examined its factorial validity for this population. Of the few studies reported, only one has been based on the original English version of the instrument (Teri, 1982a). Findings from these exploratory factor analyses (EFAs) have been diverse, with Shek (1990) reporting a 2-factor solution for a Chinese version, Baron and Laplante (1984) reporting a 3-factor solution for a

French-Canadian version, and Teri (1982a), and Larsson and Melin (1990) reporting a 4-factor solution for English and Swedish versions of the BDI, respectively.

Despite the inconsistency of factorial solutions, these EFA studies shared two common features. First, all were based on principal components analysis. Now well documented, however, are several deficiencies associated with this procedure when used as a factor analytic strategy (e.g., Borgatta, Kercher & Stull, 1986; Gorsuch, 1990; Hubbard & Allen, 1987; Snook & Gorsuch, 1989). Second, all studies used varimax rotation to achieve simple structure thereby assuming the independence of multiple factors in explaining the covariation of symptoms. Yet Tanaka and Huba (1984), in the first study to test for this assumed orthogonality, revealed these factors to be highly correlated. Such findings often suggest the presence of a 2nd-order general factor.

Byrne and Baron (1993a) addressed these limitations in their initial validation of the BDI based on three independent samples of nonclinical adolescents. As a preliminary step in their cross-validated confirmatory factor analytic (CFA) study, Byrne and Baron conducted an EFA using maximum likelihood extraction with oblique rotation on data representing their Sample 1. Findings revealed these data to be best described by a 3-factor solution[1], with factor intercorrelations ranging from .25 to .55 ($\underline{M}$=.39).

This substantial overlap in factor variances, suggested that

a general factor of depression was accounting for the three "1st-order factors" which Byrne and Baron (1993a) labelled Negative Attitude, Performance Difficulty, and Somatic Elements. To test this possibility, they proceeded next to reparameterize and test their EFA structure first as a CFA 1st-order model, and then as a 2nd-order model for data representing their second independent sample. Results revealed a 2nd-order 3-factor structure to most appropriately and parsimoniously describe the data. In a cross-validation of these findings, this final CFA model was tested and found to be totally invariant across the third independent sample of nonclinical adolescents; it is portrayed schematically in Figure 1, and represents the hypothesized model to be tested in the present study.

Insert Figure 1 about here

As depicted in Figure 1, the model under study represents a covariance structure model. Because some readers may be unfamiliar with the symbol convention associated with such models, a brief explanation is perhaps in order. Covariance structure models can be decomposed into two submodels -- a **structural model**, and a **measurement model**. The **structural model** defines the pattern of relations among unobserved hypothetical constructs. In the CFA of an assessment instrument, for example, the **structural model** formulates the pattern of factor

intercorrelations as defined by the theory upon which the instrument is based. Typically, the structural submodel is identified schematically by the presence of interrelated circles, each of which represents an hypothetical construct (or factor). Turning to Figure 1, we see an hierarchical ordering of circles such that if the page were turned sideways, the "Depression" circle would be on top, with the three smaller circles beneath it. We interpret this schema as representing one 2nd-order factor (Depression), and three 1st-order factors (Negative Attitude; Performance Difficulty; Somatic Elements). The single-headed arrows leading from the higher-order factor to each of the lower-order factors are regression paths that indicate the causal impact of Depression on the Negative Attitude, Performance Difficulty, and Somatic Elements factors; they represent the 2nd-order factor loadings. Finally, the angled arrow leading to each 1st-order factor represents residual error in their prediction from the higher-order factor of Depression.

The **measurement model** defines relations between observed variables and unobserved hypothetical constructs. In other words, it provides the link between item scores on the assessment instrument and the underlying factors they were designed to measure. The **measurement model**, then, specifies the pattern by which each item loads onto a particular factor. This submodel can be identified by the presence of rectangular boxes, each of which represents one BDI item, as indicated by the number shown. The

single-headed arrows leading from each 1st-order factor to the boxes are regression paths that link each of the factors to their respective set of observed scores; these coefficients represent the 1st-order factor loadings. For example, Figure 1 postulates that Items 1, 2, 3, etc. load onto the Negative Attitude factor. Finally, the single-headed arrow pointing to each box represents observed measurement error associated with the item variables.

One important omission in Figure 1 is the presence of double-headed arrows (↕'s) among the 1st-order factors thereby indicating their intercorrelation. This is because in 2nd-order factor analysis, all covariation among the 1st-order factors is considered to be explained by the 2nd-order factor.

The purpose of the present study, then, was to test the extent to which the model depicted in Figure 1 replicates cross-culturally for the same nonclinical adolescent population. Specifically, the study (a) tested for the validity of a 2nd-order BDI factorial structure as parameterized in Figure 1, for three independent samples of Swedish high school adolescents, (b) upon evidence of model misfit, determined the best-fitting model for each sample, and (c) tested for the invariance of factorial measurement and structure across the three samples.

Method

Sample and Procedures

Data for the present study were partly derived from an earlier study by Larsson and Melin (1990), and comprised BDI

responses from three independent samples of Swedish high school adolescents whose ages ranged from 13 to 18 years. Group 1 (n=682) was representative of a large metropolitan centre (Uppsala; population=112,000), Group 2 (n=262) of a small urban centre (Gavle; population=63,000), and Group 3 (n=200) of the rural countryside. Listwise deletion of data that were missing completely at random (see Muthen, Kaplan, & Hollis, 1987) resulted in final samples described as follows: Group 1 (n=661; 338 males, 323 females; mean age=15.5 yrs), Group 2 (n=239; 117 males, 122 females; mean age=15.5 yrs), and Group 3 (n=196; 104 males, 96 females; mean age=14.9 yrs).

Following a brief description of the purpose of the study by two graduate students, the BDI was administered to the students in their classrooms during regular school hours. Responses to all items were completed anonymously.

Instrumentation

The BDI is a 21-item scale that identifies symptoms related to cognitive, behavioral, affective, and somatic components of depression. Although originally designed for use by trained interviewers, it is now most typically used as a self-report measure (Beck et al., 1988; Kearns, Cruickshaw, McGuigan, Riley, Shaw, & Snaith, 1982; Vredenburg et al., 1985). Each item comprises four statements rated from 0 to 3 in terms of intensity, and respondents are asked to report the one which most accurately describes their own feelings; high scores represent a

more intense depression.

BDI scores comprising the present data were based on a Swedish translation of the original instrument which we refer to henceforth as the BDI-SW; item content was cross-validitated using three bilingual translators. Because we believed that for Swedish adolescents, Item 21 (Libido Loss) would be perceived as somewhat controversial, the content of this item was modified to read "Interest in Boys/Girls"; all other items remained intact.

Based on various translated versions of the BDI, estimates of internal consistency reliability (coefficient α) for nonclinical adolescents have ranged from .80 to .90 (mean α =.86; Baron & Laplante, 1984; Barrera & Garrison-Jones, 1988; Shek, 1990; Teri, 1982a). A test-retest reliability of .74 (8 weeks) has been reported for the French-Canadian version of the instrument (Baron & Laplante, 1984). Finally, Barrera and Garrison-Jones have demonstrated evidence of convergent validity with the Child Assessment Schedule (CAS; Hodges, Kline, Stern, Cytryn, & McKnew, 1982) for items measuring depression symptoms (r=.73) and with the General Self-worth subscale of the Perceived Competence Scale for Children (Harter, 1982; r=-.64); they also reported evidence of discriminant validity with CAS items measuring conduct disorder (r=.29) and anxiety (r=.29) symptoms.

Analysis of the Data

Tests for the factorial validity of the BDI-SW, and for its invariance across independent samples were based on the analysis

of covariance structures within the framework of the CFA model. Analyses were conducted in two major stages using the EQS program (Bentler, 1992). First, given the discrepant living environments from which each sample was drawn, CFA procedures were conducted for each group separately in testing the hypothesized 2nd-order factorial structure shown in Figure 1, and reported by Byrne and Baron (1993a).[2] Presented with findings of inadequate fit, we then examined parameters identified by the Lagrange Multiplier Test (LM-Test) as those that would contribute most to a significantly better-fitting model. If the inclusion of these parameters were deemed to be substantively and psychometrically reasonable, the model was respecified accordingly. Second, bearing in mind the issue of partial measurement invariance (Byrne, Shavelson, & Muthen, 1989), the final best-fitting models from Stage 1 were tested for their invariance across the three independent samples of adolescents.

Multiple criteria were used in the assessment of model fit; these were: (a) the χ^2 likelihood ratio statistic, (b) the Satorra-Bentler Scaled Statistic (S-Bχ^2; Satorra & Bentler, 1988), and (c) the Comparative Fit Index (CFI; Bentler, 1990). The S-Bχ^2 incorporates a scaling correction for the χ^2 statistic when distributional assumptions are violated. Its computation takes into account the model, the estimation method, and the sample kurtosis values (Hu, Bentler, & Kano, 1992). The S-Bχ^2 has been shown to more closely approximate χ^2 than the usual test

statistic, to have robust standard errors, and to perform as well or better than the usual asymptotically distribution-free (ADF) methods generally recommended for nonnormal multivariate data (Bentler, 1992; Hu et al., 1992). The CFI is a revised version of the Bentler-Bonett (1980) normed fit index that adjusts for degrees of freedom. It ranges from zero to 1.00 and is derived from the comparison of a restricted model (i.e., one in which structure is imposed on the data) with a null model (one in which each observed variable represents a factor). The CFI provides a measure of complete covariation in the data; a value >.90 indicates a psychometrically acceptable fit to the data. The corrected CFI value (CFI*) computed from the S-Bχ^2 statistic for the null model is also reported. It is important to note, however, that the S-Bχ^2 is not yet available for multigroup analyses in the current version of the EQS program; these values are therefore reported only for the single-group analyses.

Results

The CFA model in the present study hypothesized a priori that: (a) responses to the BDI-SW could be explained by three 1st-order factors, and one 2nd-order factor of General Depression, (b) each item would have a non-zero loading on the first-order factor it was designed to measure, and zero loadings on the other two 1s-order factors, (c) error terms associated with each item would be uncorrelated, and (d) covariation among the three 1st-order factors would be explained fully by their

regression onto the 2nd-order factor (see Figure 1).

Preliminary analyses identified one multivariate outlier for Group 1; deletion of this case resulted in a sample size of 660. Although, as expected, each group demonstrated evidence of multivariate positive kurtosis, the present data were substantially more kurtotic than all Canadian samples in general (Byrne & Baron, 1993a, 1993b; Byrne et al., 1993, 1994), and the English data from which the hypothesized model was derived in particular (Byrne & Baron, 1993a). Whereas the mean kurtosis values for the Swedish samples were 8.15, 9.17, and 6.97 for Groups 1, 2, and 3 respectively, they were 3.78, 3.41, and 2.05 for the three Canadian samples used in the Byrne and Baron (1993a) study. Given that skewness and kurtosis values are zero for data that are normally distributed, albeit values ranging from -1.00 to +1.00 may be considered to be approximately so (Muthen & Kaplan, 1985), it is easy to see how far the present data deviate from these criteria.

Although nonnormality is not likely to affect the maximum likelihood estimates, it can lead to downwardly biased standard errors which result in an inflated number of statistically significant parameters (Muthén & Kaplan, 1985). Given the abnormally high degree of kurtosis associated with the present data, it was deemed critical that final assessment of statistical fit be based on the S-Bχ^2, and on its related CFI* value, both of which correct for this violation.

Stage 1: Tests of the Hypothesized Model

Group 1

For Group 1, as indicated in Table 1, goodness-of-fit for the hypothesized 2nd-order model of BDI depression was somewhat less than adequate from both a statistical (S-Bχ^2 value) and a practical (CFI* value) perspective. (For a detailed explanation of testing covariance structure models with accompanying examples, see Byrne, 1989, 1994.) A review of the LM χ^2 statistics, however, revealed that model respecification could yield a substantially better fit if the error terms associated with three pairs of items were free to covary; these error covariances involved Items 18 (appetite loss) and 19 (weight loss), Items 20 (hypochondria) and 21 (boy/girl interest), and Items 16 (insomnia) and 17 (fatigue).

Insert Table 1 about here

Since the LM χ^2 statistics associated with the above item-pairs were distinctively larger than all remaining ones, and because findings of error covariances are not unusual in the validation of assessment instruments in general, and of psychological measures in particular, we respecified the model to include the estimation of these parameters. As shown in Table 1, this reparameterization resulted in a substantially better-fitting, and quite adequate model; all three error covariances

were statistically significant. Although a review of the LM test statistics indicated that still further model improvement could be attained by estimating several additional error covariances, these error parameters tend to be associated with the idiosyncratic interpretation of item content and thus reflect "noise" in the data, rather than any sound structural change. In the interest of parsimony, then, we considered it prudent to cease further post hoc model-fitting.

Group 2

As shown in Table 1, the initially hypothesized model for this group was also found to be inadequate. One major difference between the misfit of the model for this group compared with that of Group 1, however, was the source of the misfit as indicated by the LM χ^2 statistics. Whereas misspecification for Group 1 related to error covariances, the problem here involved the misspecification of one factor loading, in addition to several error covariances. Specifically, the LM statistics indicated that a model which allowed Item 5 (guilt) to load onto Factor 4 (Depression) would yield substantial improvement in model fit. Of course, such parameterization is not psychometrically feasible and is indicative of possible misspecification at the 1st-order level of the model. In order to test this possibility, the hypothesized model was reparameterized as a 1st-order CFA model and reestimated.

Results from this specification revealed that, as expected,

Item 5 was problematic; the LM χ^2 statistic indicated substantial model improvement if this item were permitted to cross-load onto Factor 2 (Performance Difficulty). Additionally, LM statistics associated with five error covariances were well-demarcated from those remaining, thereby indicating that the incorporation of these parameters into a subsequent model would further improve model fit.

Taking these improvement statistics into account, then, a final 2nd-order model was estimated for Group 2 in which respecifications included (a) the cross-loading of Item 5 onto Factor 2, and (b) five error covariances as follows: Item 12 (withdrawal) with Items 7 (self-dislike), 9 (suicidal), 13 (indecisiveness), 19 (weight loss), and Item 19 (weight loss) with Item 18 (appetite loss). As summarized in Table 1, this reparameterization yielded a statistically better-fitting and acceptable model.

Group 3

Consistent with findings for the two previous groups, the initially hypothesized model for Group 3, as shown in Table 1, was also inadequately specified. A review of the LM modification indices for this group indicated that the specification of three error covariances would generate a better-fitting model. Statistics for these three parameters, once again, were easily distinguishable from those remaining, and represented error covariances linking Item 21 (boy/girl interest) with Items 12

(withdrawal) and 20 (hypochondria), and between Item 16 (insomnia) and 17 (fatigue). Reestimation of a model that included these parameters led to a statistically better-fitting and adequate model (see Table 1).

A summary of the final model estimates for Groups 1, 2, and 3 is presented schematically in Figures 2-4.[3], respectively. Given the extent of nonnormality in the data, the unstandardized estimates are reported since the standardized solution does not take this distributional anomaly into account; asterisked loadings denote a parameter fixed to 1.0 for purposes of statistical identification. Parenthesized values represent z-scores adjusted for multivariate kurtosis according to the Satorra-Bentler (1988) correction formula.

Insert Figures 2, 3, and 4 about here

Reviewing these estimates, we see that while all factor loadings were statistically significant for Group 1, this was not the case for the remaining groups. Item 21 (boy/girl interest) was not significant for either Group 2 or Group 3; Item 19 (weight loss) was not significant for Group 2. Additionally, the fact that the error covariances between Items 9 (suicidal) and 12 (withdrawal) for Group 2, and between Items 20 (hypochondria) and 21 (boy/girl interest) for Group 3 were included in the model albeit they were not significant requires some explanation. In

both instances, these parameters were highly significant when based on the normal theory test-statistic. Furthermore, for each group, the model was reestimated with the related error parameter deleted; in each case this resulted in a highly significant and very substantial increase in both the uncorrected χ^2 and S-B χ^2 statistics, as well as an extremely high LM χ^2 statistic. These findings argued for the retainment of both parameters in their respective models, despite their nonsignificance based on the corrected z-statistic.

Stage 2: Tests for Invariance Across Groups

Although we know at this point that model differences exist between Swedish and Canadian adolescents, and among the Swedish groups themselves, these differences are very minor and, except for Group 2, do not involve major structural parameters. Thus, it is of interest now to test for the equality of basic factorial structure across the three independent samples under study here. It is important to note that in doing so, only parameters that are commonly specified across groups are constrained equal; all others are freely estimated. (For an elaboration of both the issue and the technique of model-testing with partial measurement invariance, see Byrne et al., 1989.)

In testing for invariant factorial structure, all 1st- and 2nd-order factor loadings were constrained equal across groups, and then tested statistically in a simultaneous analysis of the data; error covariances were allowed to be freely estimated.

Judgment of replicability was based on two criteria: (a) goodness-of-fit of the constrained model, and (b) probability level of the equality constraints as determined by the LM-Test (equality constraints with $p<.05$ being untenable). Goodness-of-fit for this 3-group constrained model yielded a $\chi^2_{(590)}=1443.65$, and a CFI=.87. However, readers are reminded that χ^2 values for the multigroup models are based on the uncorrected, rather than on the corrected (S-Bχ^2) statistic. Given the known sample size dependency of the χ^2 statistic, and the fact that it has not been scaled to correct for multivariate kurtosis, the χ^2 values are expected to be substantially larger than would be the case for the S-Bχ^2 statistic.

Examination of the probability values revealed nine 1st-order and one 2nd-order factor loadings to be nonequivalent across groups. Of these, eight represented parameter estimates that, for group 2, were significantly different from estimates for the same parameters for Groups 1 and 3; these involved the loading of Items 3 (failure), 7 (self-dislike), 8 (self-accusation), 9 (suicidal), and 14 (self-image on Factor 1, Item 11 (irritability) on Factor 2, Item 21 (boy/girl interest) on Factor 3, and Factor 1 (Negative Attitude) on Factor 4 (Depression). The two remaining inequalities related to estimated loadings for Items 13 (indecisiveness) and 20 (hypochondria) on Factor 2 for Group 3, which were significantly different from estimates for the same parameters for Groups 1 and 2. (Further

explanation and application of invariance-testing procedures are provided in Byrne 1989, 1994.)

Discussion

Findings from the present study offer strong support for the 2nd-order factorial structure of the BDI proposed by Byrne and Baron (1993a) for Canadian nonclinical adolescents. With the exception of one cross-loading for Sample 2, representative of adolescents from the small urban centre of Gavle, the hierarchical factor pattern replicated across the three Swedish samples. Although the best-fitting factor model for Swedish adolescents included correlated errors for particular pairs of items, such findings are not unexpected in multigroup analyses in general, and in cross-cultural analyses in particular. These parameters are typically unstable, and represent systematic, rather than random measurement error in item responses that may reflect bias such as yea/nay-saying and social desirability (Aish & Joreskog, 1990), or such as idiosyncratic interpretation of item content. We believe the latter to be the case here in explaining this discrepant specification between Canadian and Swedish adolescents. More specifically, we suspect that the differential interpretation of item content may be the consequence of influential cultural factors and/or the translation of the BDI from English into Swedish. Of major import, however, is that these parameters represent a trivial difference in model specification, and in no way detract from the

otherwise replicated factorial structure of the BDI for nonclinical adolescents.

Beyond this major cross-cultural finding, however, is the question of why model parameters should operate so differently for Swedish students from Gavle, compared with students from both Uppsala, and from the countryside. At first blush, one might be quick to suggest the possibility of socioeconomic factors in accounting for group differences. However, although data collection did not include information related to socioeconomic status per se, a conscious effort was made to randomly sample adolescents from a wide spectrum of schools and districts in order to address this issue. One might also point to the Swedish school system as a possible source of explanation. While nearly all adolescents attend the Swedish secondary schools (ages 13-16 years), only approximately 85% attend the high schools (ages 16-18 years); very few private schools exist in Sweden, and none exist in Uppsala and Gavle. Given that our samples were drawn from the regular school system, we expect that they are reasonably representative of the typical adolescent school population. In essence, then, we are at a loss to explain the somewhat discrepant findings associated with Sample 2, and conclude that the differential factor loading, additional error correlations and noninvariant factor loadings for the Gavle sample are likely idiosyncratic to the present data. Nonetheless, as a check of our inference here, we suggest that future research

examining depressive symptomatology in Swedish adolescents make an effort to control for the potential influence of geographic location and socioeconomic factors.

Although, in general, the finding of correlated measurement errors may be regarded as little cause for concern, the replication of correlated error terms between Items 18 (appetite loss) and 19 (weight loss) across the two urban samples (Samples 1 and 2) is nonetheless interesting and perhaps worthy of comment. This overlapping error variance can be indicative of perceived redundancy or commonality in the interpretation of item content. In this instance, it may be linked to body image, a phenomenon which seems to be an integral part of adolescents' depressive experience. Indeed, Teri (1982b) was the first to report on the relation between body image and depressive symptomatology in American adolescents; similar observations have since been reported for this population (Hops, Lewinsohn, Andrews, & Roberts, 1990; Reinherz, Frost, Stewart-Berghauer, Pakiz, Kennedy & Schille, 1990), as well as for younger children (McCabe & Marwit 1993). As noted by Baron (1993), the importance of such findings lies with the perception and evaluation by adolescents of their own physical self, a component of self-esteem, and one of the central elements of depression. Indeed, further endorsement of this perspective has been provided by Allgood-Merten, Lewinsohn, and Hops (1990) who proposed that global self-esteem and the more specific body-esteem be integral

components of any model hypothesized as representing adolescent depression. Consonant with previous research, then, our results suggest that the interplay between self-esteem and body-esteem holds also for Swedish adolescents' in their depressive experiences.

Finally an issue that needs to be addressed is the severely kurtotic distribution of most BDI-SW item scores for the Swedish samples. Although BDI scores for nonclinical Canadian adolescents have been found also to be leptokurtic (i.e., positively kurtotic), those for Swedish adolescents are substantially more so. The puzzling question, of course, is why this should be the case. One possible explanation may lie with the issue of gender. For example, in a recent investigation of the factorial equivalence of the BDI across Swedish nonclinical male and female adolescents, Byrne, Baron, Larsson, and Melin (in press) found scores for females to be substantially more kurtotic than those for males. They suggested that socialization factors integral to Swedish society at large likely contributed to this differential response pattern. Although admittedly speculative at this point, their findings call for future research that more closely examines gender differences as they bear on adolescent depressive experience in Scandinavian countries generally, and in Sweden specifically.

Although our discussion has focused largely on differential psychometric and sampling distributional features related to the

present data, of substantially more import is the fact that the 2nd-order BDI factor structure proposed by Byrne and Baron (1993a) was adequately successful in describing these data for Swedish nonclinical adolescents. These findings have important implications for Swedish clinicians in their identification of depressive symptoms in this population of young people. Beyond these national considerations, however, our findings add importantly to a growing body of evidence that argues for an hierarchical factorial structure of the BDI to most appropriately model depressive symptoms in nonclinical adolescents. From a practical, as well as psychometric perspective, it now seems imperative that construct validity research related to the BDI establish whether this same hierarchical structure holds for clinically depressed adolescents.

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Footnotes

1. Although there was a hint of a possible 4-factor solution, the pattern was marred by numerous cross-loadings (>.20) that obviated the attainment of any simple structure worthy of meaningful interpretation.
2. Had the samples been drawn from the same urban (or rural) population, the hypothesized model would have been tested on on one group only (say, Group 1) which would have served as a calibration group. The final model based on this group would then have been validated across the other two groups.
3. Item 21 was altered to reflect the change in content for the BDI-SW.
4. It can also occur as a consequence of redundant item content (i.e., multiple items that elicit essentially the same information). This would not appear to be the case with the BDI.

Table 1

Summary of Fit Statistics for 2nd-order Models of BDI Factorial Structure

Model	S-Bχ^2	df	CFI*
Group 1			
Hypothesized Model	376.77	187	.82
Final Model	301.14	184	.90
• 3 correlated errors			
Group 2			
Hypothesized Model	281.08	187	.81
Final Model	219.69	181	.92
• 1 cross-loading			
• 5 correlated errors			
Group 3			
Hypothesized Model	254.21	187	.82
Final Model	224.79	184	.90
• 3 correlated errors			

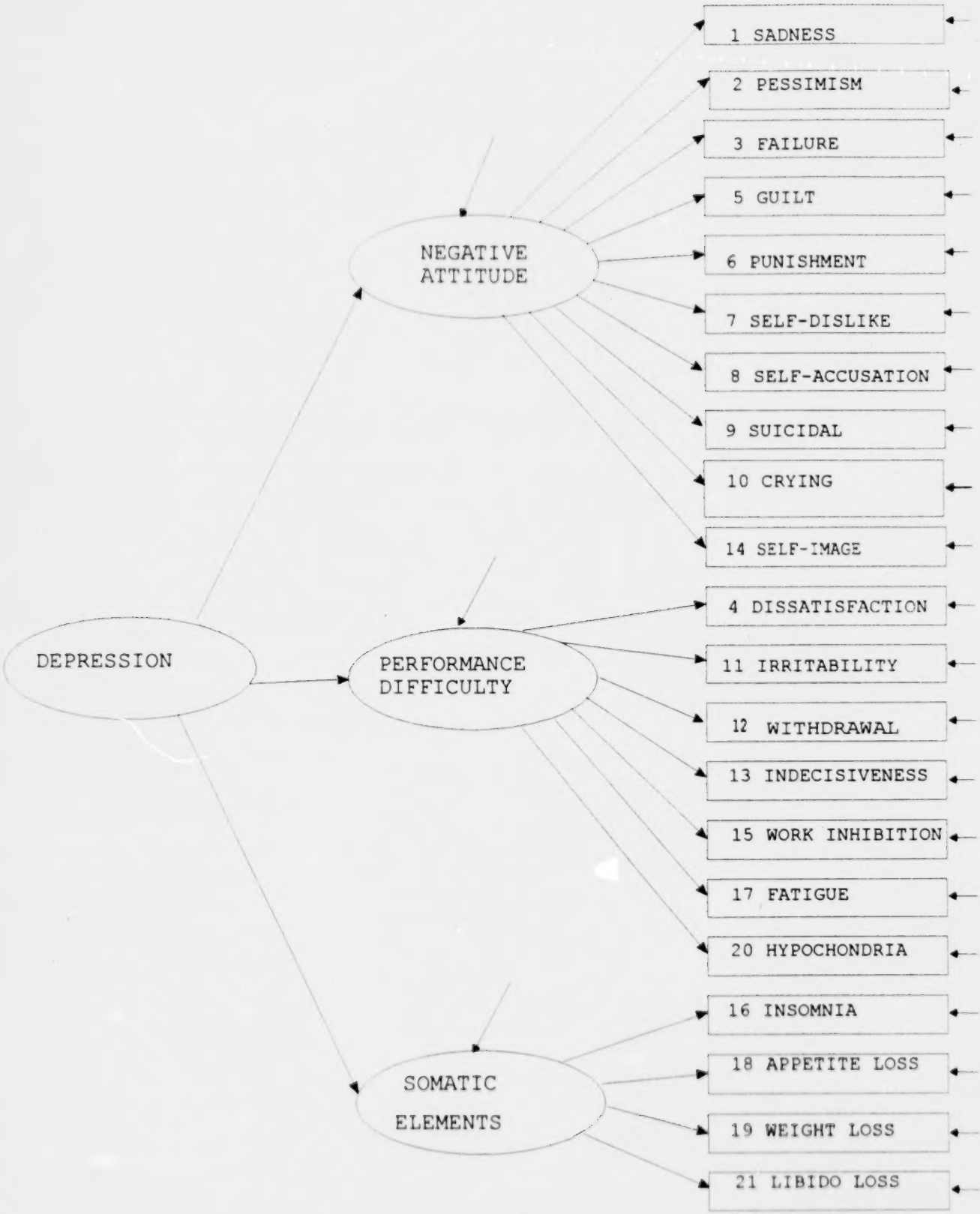
Figure Captions

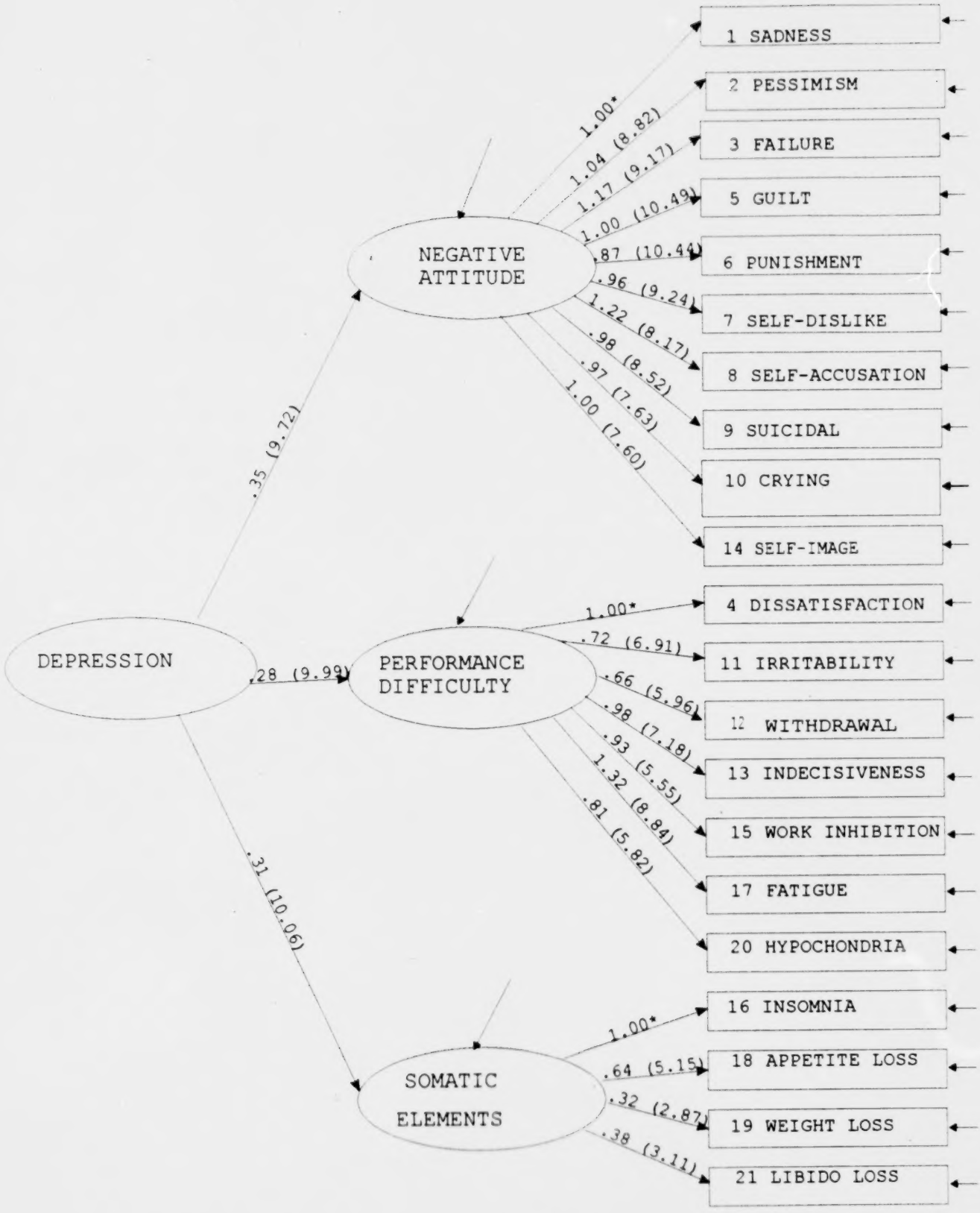
Figure 1. Hypothesized Model of Factorial Structure of the Beck Depression Inventory Based on a Cross-validated Study of Canadian Adolescents (Byrne & Baron, 1993a). Copyright permission to be obtained from the American Association of Counseling and Development.

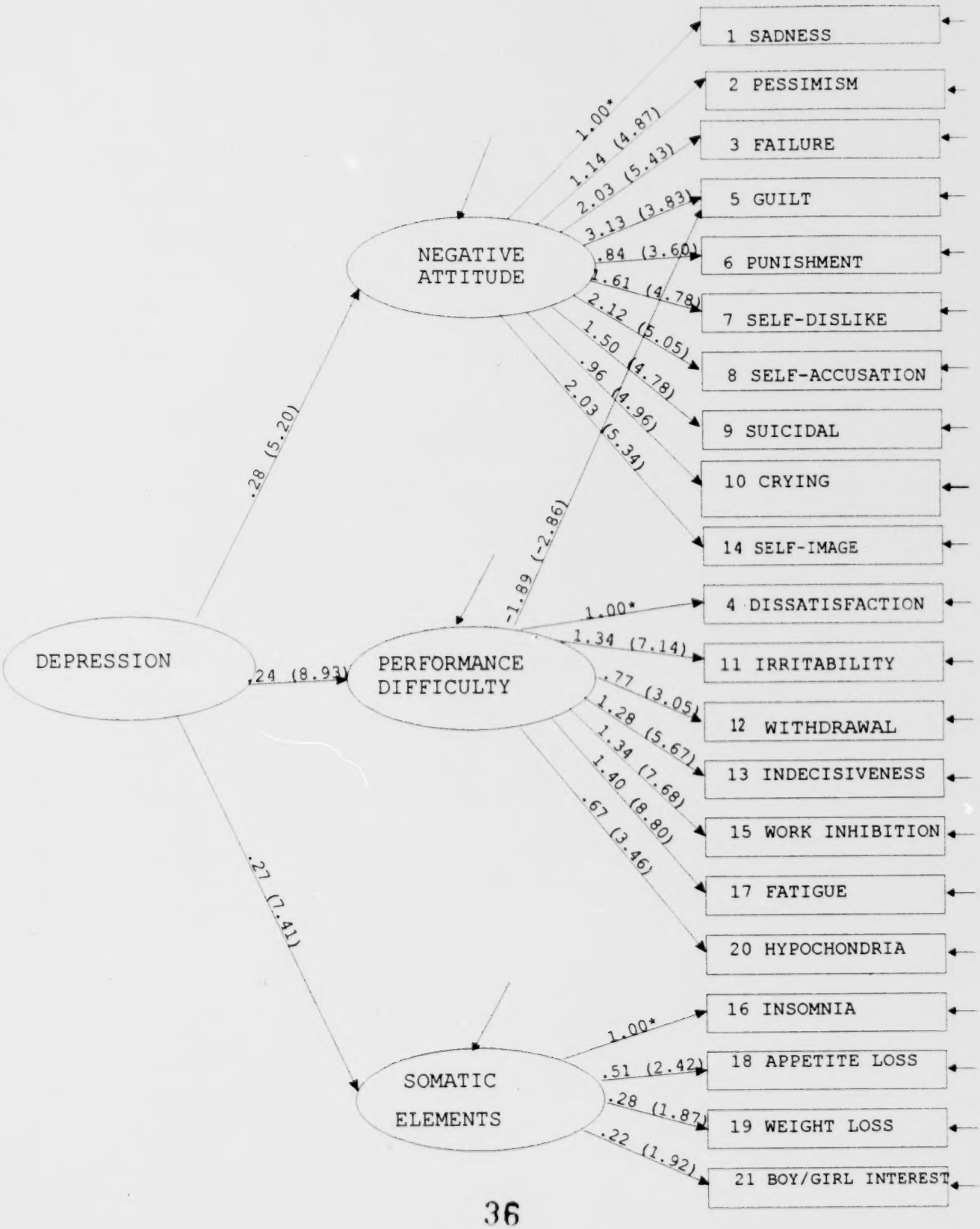
Figure 2. Final Model of Factorial Structure of the Beck Depression Inventory for Swedish Adolescent Sample 1 (Uppsala). Correlated error estimates were: Items 16/17 = .09 (4.91); Items 20/21 = .04 (2.11); Items 18/19 = .06 (4.50).

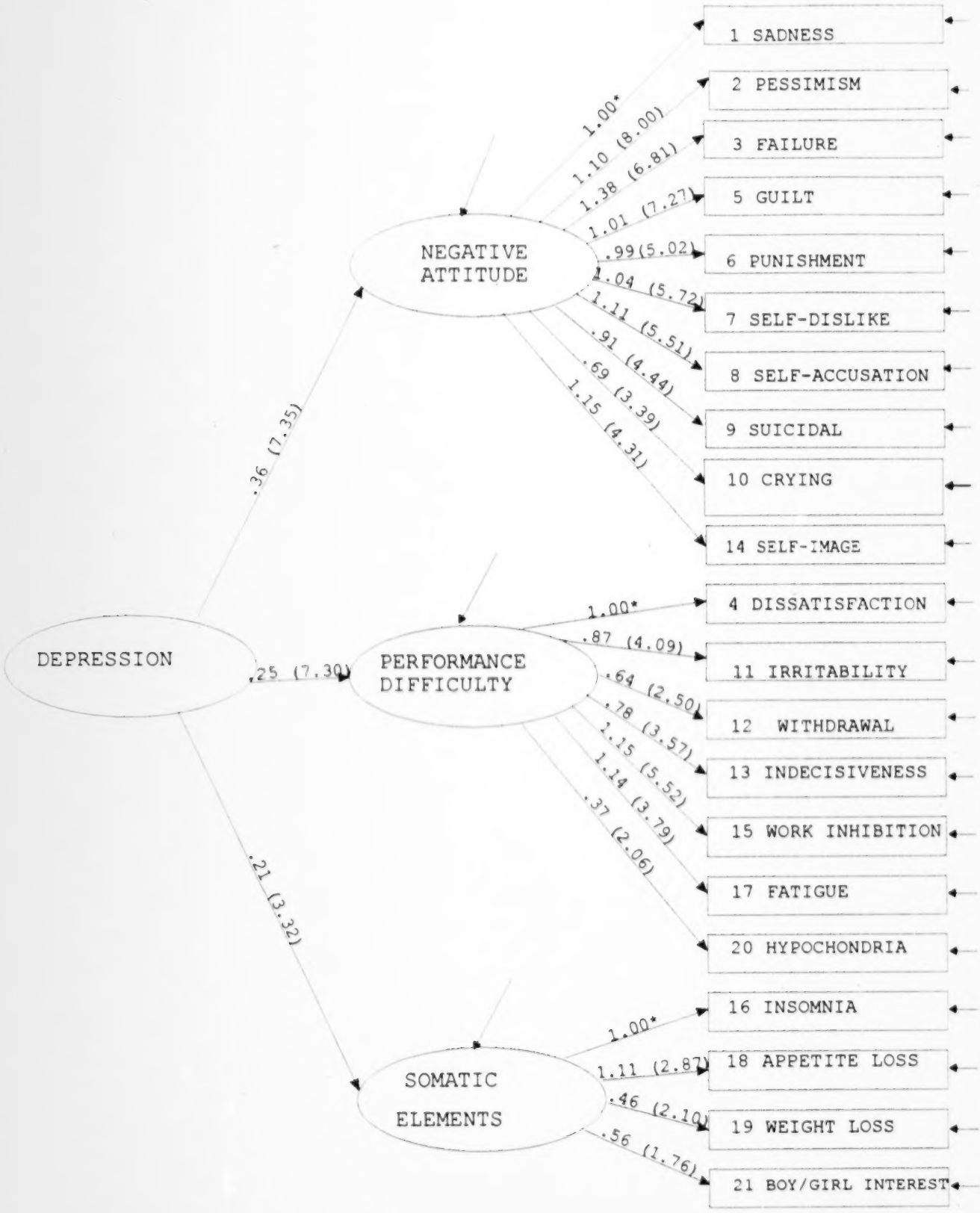
Figure 3. Final Model of Factorial Structure of the Beck Depression Inventory for Swedish Adolescent Sample 2 (Gavle). Correlated error estimates were: Items 7/12 = .04 (2.13); Items 9/12 = .06 (1.43); Items 12/13 = .04 (2.47); Items 12/19 = .03 (2.79); Items 18/19 = .05 (2.41).

Figure 4. Final Model of Factorial Structure of the Beck Depression Inventory for Swedish Adolescent Sample 3 (Countryside). Correlatederror estimates were: Items 12/21 = .06 (2.25); Items 16/17 = .13 (2.71); Items 20/21 = .04 (1.61).


DEPRESSION
NEGATIVE ATTITUDE
PERFORMANCE DIFFICULTY
SOMATIC ELEMENTS
1 SADNESS
2 PESSIMISM
3 FAILURE
5 GUILT
6 PUNISHMENT
7 SELF-DISLIKE
8 SELF-ACCUSATION
9 SUICIDAL
10 CRYING
14 SELF-IMAGE
4 DISSATISFACTION
11 IRRITABILITY
12 WITHDRAWAL
13 INDECISIVENESS
15 WORK INHIBITION
17 FATIGUE
20 HYPOCHONDRIA
16 INSOMNIA
18 APPETITE LOSS
19 WEIGHT LOSS
21 LIBIDO LOSS


DEPRESSION
NEGATIVE ATTITUDE
PERFORMANCE DIFFICULTY
SOMATIC ELEMENTS
.35 (9.72)
.28 (9.99)
.31 (10.06)
1.00*
1.04 (8.82)
1.17 (9.17)
1.00 (10.49)
.87 (10.44)
.96 (9.24)
1.22 (8.17)
.98 (8.52)
.97 (7.63)
1.00 (7.60)
1 SADNESS
2 PESSIMISM
3 FAILURE
5 GUILT
6 PUNISHMENT
7 SELF-DISLIKE
8 SELF-ACCUSATION
9 SUICIDAL
10 CRYING
14 SELF-IMAGE
1.00*
.72 (6.91)
.66 (5.96)
.98 (7.18)
.93 (5.55)
1.32 (8.84)
.81 (5.82)
4 DISSATISFACTION
11 IRRITABILITY
12 WITHDRAWAL
13 INDECISIVENESS
15 WORK INHIBITION
17 FATIGUE
20 HYPOCHONDRIA
1.00*
.64 (5.15)
.32 (2.87)
.38 (3.11)
16 INSOMNIA
18 APPETITE LOSS
19 WEIGHT LOSS
21 LIBIDO LOSS


1 SADNESS
2 PESSIMISM
3 FAILURE
5 GUILT
6 PUNISHMENT
7 SELF-DISLIKE
8 SELF-ACCUSATION
9 SUICIDAL
10 CRYING
14 SELF-IMAGE
4 DISSATISFACTION
11 IRRITABILITY
12 WITHDRAWAL
13 INDECISIVENESS
15 WORK INHIBITION
17 FATIGUE
20 HYPOCHONDRIA
16 INSOMNIA
18 APPETITE LOSS
19 WEIGHT LOSS
21 BOY/GIRL INTEREST
NEGATIVE ATTITUDE
PERFORMANCE DIFFICULTY
SOMATIC ELEMENTS
DEPRESSION
1.00*
1.14 (4.87)
2.03 (5.43)
3.13 (3.83)
.84 (3.60)
1.61 (4.78)
2.12 (5.05)
1.50 (4.78)
.96 (4.96)
2.03 (5.34)
-1.89 (-2.86)
1.00*
1.34 (7.14)
.77 (3.05)
1.28 (5.67)
1.34 (7.68)
1.40 (8.80)
.67 (3.46)
1.00*
.51 (2.42)
.28 (1.87)
.22 (1.92)
.28 (5.20)
.24 (8.93)
.27 (7.41)



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